Limited Dependence of the Human Sex Ratio on Birth Order and Parental Ages

Michael S. Teitelbaum,[1] Nathan Mantel,[2] and Charles R. Stark[3]

INTRODUCTION

The voluminous literature on the ratio of male to female human births suggests that well over 30 factors exert an influence on it [1–3], but a clearly significant role has been demonstrated for only a few of these. To some extent, what may be lacking is adequate confirmatory study of suspect factors and, relating to this, adequate methodology for studying certain factors while removing the confounding effects of others. A further difficulty derives from the unavailability of data which permit controlling for more than one or two of the suggested factors at once. Data which do provide such control are usually relatively small in size, and, as will become evident later, study sizes necessary for sex-ratio studies may be quite large.

The relationships between the sex ratio and three of these more than 30 factors—birth order, paternal age, and maternal age—have been very extensively studied. Wicksell [4, 5] found contradictory evidence in data for Berlin. Ewart [6] found no significant relationship between sex ratio and maternal age. Russell [7], using American data, found a significant negative effect of paternal age and no significant effect of maternal age. Ciocco [8], also using American data, concluded that there was a negative effect of the combined parental ages on sex ratio but was unable to separate the parental age effects from one another or from that of birth order. Lejeune and Turpin [9], using similar data, came to the same conclusion but were also unable to separate maternal age, paternal age, and birth-order effects satisfactorily.

Novitski [10] concluded that paternal age had a negative effect on sex ratio and that maternal age had no significant effect. His analysis was later criticized by Bernstein [11] on the grounds that he made no correction for the correlation between maternal and paternal ages. Novitski and Sandler [12] extended the Novitski analysis, using a multiple linear-regression analysis, and found a significant negative effect of paternal age and no significant effect of maternal age. In a separate analysis, they also demonstrated a significant negative relationship between sex ratio and birth

Received July 7, 1970; revised November 24, 1970.

[1] Office of Population Research, Princeton University, Princeton, New Jersey 08540; and Epidemiology Branch, National Cancer Institute, National Institutes of Health, Bethesda, Maryland 20014.

[2] Biometry Branch, National Cancer Institute, National Institutes of Health, Bethesda, Maryland 20014.

[3] Epidemiology and Biometry Branch, National Institute of Child Health and Human Development, National Institutes of Health, Bethesda, Maryland 20014.

order. Unfortunately, they were unable to provide control for the effects of birth order in their analysis of maternal and paternal age effects, and vice versa. Since these three variables are highly intercorrelated, Novitski and Sandler's findings cannot be considered conclusive. Recently, Moran, Novitski, and Novitski [13] acknowledged some of the difficulties of the Novitski-Sandler report. However, they too are unable to provide simultaneous control of birth order, paternal age, and maternal age, and their reanalysis of the Novitski-Sandler data leads them to the same findings.

In another recent report, Pollard [14] has attempted to analyze the effects of birth order, paternal age, and maternal age on the sex ratio at birth in Australia. Once again, he was unable to provide simultaneous control for all three factors, and his findings must therefore be viewed with caution.

Apparently the only available data providing simultaneous tabulations of sex by birth order, paternal age, and maternal age are those reported by Novitski and Kimball [15]. While these data are clearly the best available, Novitski and Kimball's analysis is not amenable to clear interpretation. In the present paper, we present a reevaluation of the Novitski-Kimball data in order to show the consequences of studying separately each of the three factors (birth order, paternal age, and maternal age) while simultaneously providing adequate control for the others.

THE PROBLEM

In 1958, Novitski and Kimball reported their findings on the effects of birth order and parental ages on the sex ratio at birth. Their procedure for removing confounding effects consisted of fitting a quadratic regression surface to data on the male proportion of live births and retaining only significant regression variables. Their data, obtained from a special tabulation by the U.S. National Office of Vital Statistics, are unique in that they are classified simultaneously by paternal age, maternal age, and live-birth order (but apparently not for race) for live births occurring in the United States in 1955. These births represent a population size of 3,645,750 for which all the relevant information was known, but with parental ages restricted to 49 years and under for fathers and 39 years and under for mothers. The data are presented in detailed form in table 1 of Novitski and Kimball [15].

In their final regression fit to these data, Novitski and Kimball found a negative effect for increasing birth order and for the square of birth order,* a small negative effect for increasing paternal age, and a moderately large positive effect for birth order × paternal age (interaction). This last positive effect showed the largest t ratio in their findings, but it is not clear how this should be interpreted.† Because

* From the text, there would seem to be a typographical error in table 4 of Novitski and Kimball. We are assuming their b_{23} (maternal age × birth order) value in that table to represent b_{33} (square of birth order).

† Care needs to be taken in interpreting the significance of individual regression coefficients in Novitski and Kimball. Because of the high correlation between birth order and square of birth order over the range studied, the high significance of the two together shows up as only limited significance for each individually. Actually, there is some question about their practice of deleting the linear term when it does not give significantly improved fit over using quadratic terms only. The reverse kind of deletion, however, is proper. The problem arises when the origin of the independent variable or variables is arbitrary. At one location of the origin, the linear term may be significant and so retained,

of this interaction effect, the contours of their figure 1 show that, if paternal age were high enough, the birth-order effect could be positive rather than negative; and, if birth order were high enough, the paternal age effect could also be positive rather than negative. These contours are reproduced in figure 1 of the present paper, together with supporting data from table 1 of Novitski and Kimball [15]. For ease of inspection, these data are expressed in code form as:

$$\frac{(\text{sex proportion} - 0.510) \times 1000}{\text{number of births in cell (in thousands)}}\,.$$

Thus a 51.5% male sex percentage arising among 25,187 births would be shown as $+5/25$. The correspondence of the fitted contours to the coded data is not obvious upon inspection.

The present report is directed to a different method for evaluating the unconfounded effects of the three study factors. The need for this reevaluation stems from indications within the Novitski-Kimball report itself that the functional form fitted may have been improper, and in such a way as to produce an artifactually significant interaction term of the kind found. The possibility of the failure of their quadratic model to fit the data on sex ratio was, in fact, recognized by Novitski and Kimball. They noted that under their weighting procedure the residual sum of squares of 218.95 with 165 df is distributed approximately as χ^2 and corresponds to a significance

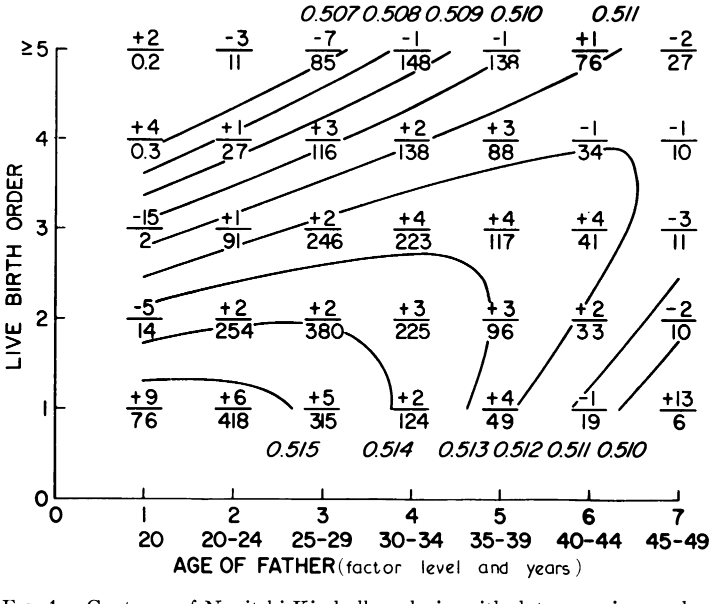

FIG. 1.—Contours of Novitski-Kimball analysis, with data superimposed

---

while at some other, the linear term is exactly zero and so can be dropped with no essential change. But at in-between origin locations, the linear term may be nonsignificant; if it is then deleted, the resulting fitted regression function becomes different. Were the origin for an independent value a true zero, then the practice used by the authors could be justified under otherwise appropriate circumstances.

probability of .003. The likely explanation is that while a quadratic surface may be used to approximate a small portion of any true surface, it cannot be expected to fit a surface in its entirety. (Alternatively, there might actually be a great excess over simple random variation, but both our own analysis and that of Novitski and Kimball do not suggest any such important departures once birth-order effects are removed.)

Just how the improper use of a quadratic surface can produce artifacts is of interest. In two dimensions, if a parabola is fitted over a region where the slope is always positive but diminishing to the right, the best-fitting parabola will be one which eventually, and perhaps soon, progresses downward. In three dimensions, if two of them represent independent variables, each with positive effect but not fully reinforcing of each other, a quadratic surface analysis may incorrectly yield a negative-interaction regression coefficient under which the fitted surface shows eventual decrease rather than only moderate increase. A fitted surface of this type occurring in the Novitski-Kimball analysis makes difficult any definite statement about the direction of the birth-order or paternal age effects. In any case, it should be remarked that the final regression equation fitted by Novitski and Kimball accounts for only a small range of variation in the sex ratio. In figure 1, it can be seen that their male sex proportion contours run from a low of 0.507 to a high of 0.515.

MATERIALS AND METHODS

The method of obtaining unconfounded effects to be employed here stems from one originally presented by Mantel and Haenszel [16] and subsequently extended by Mantel [17]. An interesting epidemiological application of the Mantel-Haenszel method is one by Stark and Mantel [18], in which the authors ascertained the separate effects of birth order and maternal age in Down's syndrome and in childhood leukemia. The method reverses the regression-type approach of Novitski and Kimball. For example, instead of being directed to how the male sex proportion changes with birth order, it is directed to how the average birth order differs for male and female children. If we are adjusting for paternal age, a comparison of birth-order averages is made for each paternal age, the separate comparisons then being combined into a summary comparison. Alternatively, adjustment can be made for maternal age alone, for both parental ages, or for neither.

Associated with each comparison of average differences between male and female children, whether of birth order or of either parental age, the Mantel-Haenszel procedure yields a continuity-corrected χ^2 with 1 df, which permits assessing the statistical significance of the observed difference. As a measure of biological significance or importance, we use the measure R, the "relative odds" as defined in table 1 of Mantel and Haenszel [16], and there used as an approximate measure of relative risk. In calculating R, some particular level of a factor must be identified as a standard. Here, for birth order, we use first-born children; for either parental age, we take children born of parents 20–24 years old. In the present context, the relative odds is a particularly appropriate measure, since it corresponds to the ratio of two sex ratios.*

* Note, however, that a 0.01 shift in the male sex proportion, from 0.50 to 0.51, results in approximately a 0.04 shift in the relative odds, which is then approximately 1.04. This is also true of the simple ($M/F \times 100$) sex ratio which is conventionally used in sex-ratio studies. An $[M/(M + F)]$ sex proportion of 0.50 is equivalent to a sex ratio of 100, while a sex proportion of 0.51 corresponds approximately to a sex ratio of 104.

Essentially, the method is directed toward the main effects of each factor considered, with the investigator free to examine the results of his analysis for any indication of interaction-like effects. When viewed in this way, the kind of analysis made does not differ importantly from a simple linear-regression analysis of, say, the male sex proportion on birth order, since the covariance of the 0-1 variable, maleness of child, with birth order of child is the same whichever is regarded as the independent variable.

The calculations necessary for the Mantel-Haenszel procedure are conceptually simple, and an uncomplicated analysis (e.g., controlling for only one factor) can be done by hand calculations. However, more complex analyses become cumbersome and time-consuming, and for this reason a generalized FORTRAN IV program has been developed. All calculations for this analysis have been carried out on the National Institutes of Health IBM 360/50 computer.

RESULTS

The summary analyses of the three study factors are shown in table 1, with each factor evaluated taking into account neither, either, or both of the remaining two factors. For each analysis made, table 1 gives the summary χ^2, its associated average-score difference, and supporting relative-odds values at various levels of the study factor when a specified level is designated as standard.

The crude (unadjusted) analysis for the birth-order effect yields a χ^2 of 53.65, which is very highly significant ($P < .00001$).* Even when adjustment is made for either or both of the other two variables, the resultant χ^2 remains similarly significant.

We must note, however, that associated with these significant χ^2s is an average birth-order score for male children only 0.008–0.010 below that for female children. Similarly, the relative odds associated with these significant χ^2s are rather close to unity, in contrast to values on the order of 1.5–10.0 found in many epidemiological investigations, for example. These relative odds do show some progressiveness in their departures from unity, but it is apparently the large amount of data (over 3.6 million births) which helps give rise to significant χ^2s despite small effects. The summary relative odds at the highest birth order is 0.026 below that at the first, indicating only a small decline in the sex ratio with increasing birth order.

The maternal and paternal age effects summarized in table 1 are even less marked than the birth-order effect. While more or less significant χ^2s for paternal age and maternal age arise in the crude analyses, in both cases indicating a decreasing sex ratio with increasing parental age, any statistical significance is lost following adjustment, and specifically when adjustment is made for birth order.

Tables 2–4 show detailed separate analyses for birth-order, paternal age, and maternal age effects, respectively. Thus, in table 2, significance is sought (via χ^2 and average birth-order scores) for a birth-order effect within paternal age categories, with maternal age alternatively taken into account or ignored; a corresponding birth-order analysis is made within maternal age categories, with paternal age adjusted for or ignored. In tables 3 and 4, similar reciprocal analyses are made for parental ages.

In Table 2, a consistent negative influence of birth order on the sex ratio shows

* All of these probability levels are for two-tailed tests, although in some instances one-tailed tests may be justified. In such cases, the given probability levels may be halved.

TABLE 1

SUMMARY ANALYSES OF BIRTH ORDER, PATERNAL AGE, AND MATERNAL AGE

	Continuity-corrected χ^2 (1 df)	Average Score Difference Males—Females	Relative Odds at Various Study Factor Levels 1	2	3	4	5	6	7
			Factor Level 1 (First Birth Order = 1.000)						
Birth-order analysis:									
Crude analysis......	53.65	−0.010	1.000	0.989	0.991	0.988	0.972	......	......
Paternal age adjustment only........	42.22	−0.008	1.000	0.990	0.993	0.993	0.974	......	......
Maternal age adjustment only........	49.11	−0.009	1.000	0.989	0.992	0.990	0.974	......	......
Both parental age adjustments......	45.26	−0.008	1.000	0.989	0.993	0.991	0.974	......	......
			Factor Level 2 (Paternal Age 20–24 = 1.000)						
Paternal age analysis:									
Crude analysis......	11.35	−0.005	1.011	1.000	0.994	0.994	0.993	0.990	0.983
Birth-order adjustment only........	0.01	+0.000	1.005	1.000	1.000	0.999	1.004	0.999	0.997
Maternal age adjustment only........	6.00	−0.002	1.008	1.000	0.996	0.990	0.996	0.992	0.998
Both birth order and maternal age.....	1.37	−0.001	1.004	1.000	0.998	0.992	0.999	0.997	1.006
			Factor Level 2 (Maternal Age 20–24 = 1.000)						
Maternal age analysis:									
Crude analysis......	5.86	−0.003	1.007	1.000	0.996	1.000	0.993	......	......
Birth-order adjustment only........	2.04	+0.001	1.001	1.000	1.000	1.009	1.001	......	......
Paternal age adjustment only........	0.01	+0.000	1.003	1.000	0.999	1.004	0.996	......	......
Both birth order and paternal age......	2.68	+0.001	0.999	1.000	1.001	1.008	1.000	......	......

up in every paternal or maternal age category, and the results are little influenced by whether the alternative parental age has been adjusted for or ignored. In most instances, this negative influence is statistically significant, and it is likely that the sparseness of available data is responsible for lack of statistical significance in a number of instances. In any case, the analysis does not show any clear pattern of a reversed or diminishing effect of birth order in higher paternal age categories as the Novitski-Kimball fit would have suggested.

The corresponding detailed analysis for paternal age given in table 3 does not show as consistent a pattern as that which occurred for birth order. The paternal age effect is sometimes positive, sometimes negative, and is nonsignificant overall, as shown in table 1 when the birth-order adjustment is made. Of interest here, however, is the significant negative effect for paternal age which occurs at birth-order 1 (though

TABLE 2

ANALYSIS OF BIRTH-ORDER EFFECT BY PATERNAL AND MATERNAL AGE CATEGORIES

Birth-Order Analysis within Paternal Age Categories	Paternal Age						
	Under 20	20–24	25–29	30–34	35–39	40–44	45–49
Adjusting for maternal age:							
χ^2 (1 df)	8.15	13.34	26.32	4.18	5.59	0.08	1.81
Average birth-order score difference (males−females)	− 0.009	− 0.007	− 0.011	− 0.006	− 0.009	− 0.002	− 0.015
Ignoring maternal age:							
χ^2 (1 df)	10.58	14.15	25.97	2.52	5.43	0.01	2.25
Average birth-order score difference (males−females)	− 0.011	− 0.008	− 0.011	− 0.004	− 0.009	− 0.001	− 0.017

Birth-Order Analysis within Maternal Age Categories	Maternal Age						
	Under 20	20–24	25–29	30–34	35–39	40–44	45–49
Adjusting for paternal age:							
χ^2 (1 df)	11.08	8.28	11.51	20.97	1.32	....	....
Average birth-order score difference (males−females)	− 0.007	− 0.006	− 0.008	− 0.015	− 0.005	....	....
Ignoring paternal age:							
χ^2 (1 df)	13.72	9.42	12.63	21.28	1.35	....	....
Average birth-order score difference (males−females)	− 0.007	− 0.006	− 0.009	− 0.015	− 0.006	....	....

TABLE 3

ANALYSES OF PATERNAL AGE EFFECT BY BIRTH-ORDER AND MATERNAL AGE CATEGORIES

Paternal Age Analysis within Birth-Order Categories	Birth Order				
	1	2	3	4	5+
Adjusting for maternal age:					
χ^2 (1 df)	6.41	2.17	0.71	2.58	6.79
Average paternal age score difference (males−females)	−0.004	−0.002	+0.002	−0.004	+0.007
Ignoring maternal age:					
χ^2 (1 df)	8.14	0.24	1.76	0.49	6.43
Average paternal age score difference (males−females)	−0.006	+0.001	+0.004	−0.003	+0.009

Paternal Age Analysis within Maternal Age Categories	Maternal Age				
	<20	20–24	25–29	30–34	35–39
Adjusting for birth order:					
χ^2 (1 df)	3.42	0.24	0.14	0.02	0.05
Average paternal age score difference (males−females)	−0.004	−0.001	−0.001	−0.000	−0.001
Ignoring birth order:					
χ^2 (1 df)	6.10	1.28	1.02	0.68	0.22
Average paternal age score difference (males−females)	−0.006	−0.002	−0.002	−0.002	−0.002

TABLE 4

ANALYSIS OF MATERNAL AGE EFFECT BY BIRTH-ORDER AND PATERNAL AGE CATEGORIES

MATERNAL AGE ANALYSIS WITHIN BIRTH-ORDER CATEGORIES	BIRTH ORDER				
	1	2	3	4	5+
Adjusting for paternal age:					
χ^2 (1 df)	0.33	7.08	0.02	2.29	2.01
Average maternal age score difference (males−females)	+0.001	+0.004	−0.000	+0.003	−0.003
Ignoring paternal age:					
χ^2 (1 df)	2.76	5.71	1.07	0.66	0.57
Average maternal age score difference (males−females)	−0.003	+0.005	+0.002	+0.002	+0.002

MATERNAL AGE ANALYSIS WITHIN PATERNAL AGE CATEGORIES	PATERNAL AGE						
	<20	20–24	25–29	30–34	35–39	40–44	45–49
Adjusting for birth order:							
χ^2 (1 df)	2.27	0.06	1.01	3.58	0.00	0.46	0.11
Average maternal age score difference (males−females)	−0.004	+0.000	+0.001	+0.003	−0.000	+0.003	−0.002
Ignoring birth order:							
χ^2 (1 df)	3.90	0.59	0.15	2.67	0.22	0.45	0.41
Average maternal age score difference (males−females)	−0.005	−0.001	−0.001	+0.003	−0.001	+0.003	−0.005

that effect is almost matched in magnitude at birth-order 4), while a significant positive effect occurs in the highest birth order. Although the pattern is not clear and a corresponding phenomenon did not show up in the detailed birth-order analysis, this does suggest that the data do contain some elements which make for the kind of interaction reported by Novitski and Kimball.

In the detailed maternal age analysis covered in table 4, there is also no consistent pattern of effects. While there is a significant positive maternal age effect for second birth-order children, this can be discounted in view of the multiplicity of significance tests arising in this analysis.

DISCUSSION

Our primary finding here is that of a significant negative effect of birth order on the human sex ratio at birth. This effect, which shows up whether or not the confounding effects of parental ages are removed, is quite limited, however, and is reflected in a reduction of only 0.026 in the relative odds of a male birth among fifth and later births as compared with those of first births. This birth-order effect apparently gives rise to negative parental age effects on the sex ratio when only the crude data are examined. Both paternal and maternal age effects disappear when birth order is taken into account. However, there is some suggestion in the data of a negative paternal age effect at the first-birth order but a positive effect in the highest birth-order group considered. This is in accord with the Novitski-Kimball analysis which revealed a significant birth-order × paternal age interaction. For reasons we have noted above, the Novitski-Kimball analysis implicitly accorded that interaction undue importance. If the interaction is taken as real, it probably reflects other confounding effects which could not be sorted out from the data. One possibility is the confounding effect of race. Ciocco [8] and Visaria [19] have argued that Negroes show a significantly lower sex ratio than do whites, and the Novitski-Kimball data represent a pooling of "white" and "colored" births. Different patterns of family formation between whites and Negroes (e.g., a higher probability for young Negroes to have many children) might therefore affect the distribution of sexes in some of the birth-order and parental age groups, and give rise to an apparent "interaction effect."

While the statistical significance of the negative birth-order effect is high, the effect's limited magnitude could indicate that it is only a reflection of some other mechanism. A candidate mechanism is one in which families differ in their intrinsic sex ratios; if some couples are motivated to continue family formation until a male child is born, there could be higher proportions of families with low sex ratios at the higher birth orders [20]. Another attractive mechanism would be an immunological interaction between the mother and the conceptus. There is extensive literature on this possibility, but the matter is as yet unresolved [1, 21].

Whatever mechanism may be operating, it is clear that its net influence must be rather small and therefore would be difficult to detect in a study of moderate size. The present results suggest that studies directed to detecting effects of the magnitude found here must be large indeed. Here, we have obtained χ^2s of about 50 with data available for about 3.6 million children. Probably the birth-order effect would have been detected, but with less clearly established significance, using a study size one-fifth to one-sixth as large, say, 500,000–600,000 children. Studies any smaller than this are likely to prove unrewarding unless it can be anticipated that they are directed toward factors which have greater effects on the sex ratio than has birth order. In some instances, sex-ratio data are available at little or no cost, as was the case for the present analysis. However, data for more general studies of factors influencing the sex ratio cannot be expected to be so freely accessible. Thus, investigators must either mount extravagant studies or display sharp insight into what factors can have relatively important effects, that is, ones that produce a change of 5% or more in the male sex percentage.

SUMMARY

Reanalysis of the data obtained by Novitski and Kimball [15] does not support some of their major conclusions. Novitski and Kimball, using a multiple quadratic regression surface analysis to disentangle the confounding effects of birth order, paternal age, and maternal age on sex ratio, reported the following: a significant negative effect for birth order and birth order squared, a small significant negative effect for paternal age, and a moderately large positive effect for birth-order × paternal age interaction.

Using the extended Mantel-Haenszel χ^2 procedure on the data presented by Novitski and Kimball, the present analysis can confirm the presence of a significant negative birth-order effect but cannot confirm significance for the effect of paternal

age. There are also indications of a reversal of the birth-order effect as paternal age increases. While this suggests that the data do contain elements of the kind of interaction reported by Novitski and Kimball, the functional form of their analysis may tend to overstate the importance of this effect. In any case, it is possible that the confounding effect of race may be affecting the analysis of these data.

REFERENCES

1. EDWARDS AWF: Genetics and the human sex ratio. *Advances Genet* 11:239–272, 1962
2. STEVENSON AC, BOBROW M: Determinants of sex proportions in man, with consideration of the evidence concerning a contribution from X-linked mutations to intrauterine death. *J Med Genet* 4:190–221, 1967
3. TEITELBAUM MS: Factors affecting the sex ratio in large populations. *J Biosocial Sci*, suppl. 2, 1970, pp 61–71
4. WICKSELL SD: Sex proportion and parental age. *Lunds Universitets Arsskrift*, N. F., Avd. 2, vol 22, no. 6, 1926
5. WICKSELL SD: *Kungliga Fysiografiska Sallskapets Handlingar*. N. F., vol 37, no. 6, 1926
6. EWART RJ: The influence of age of the grandparent at the birth of the parent on the number of children born and their sex. *J Hyg* 15:127–162, 1915
7. RUSSELL WT: Statistical study of the sex ratio at birth. *J Hyg* 36:381–401, 1936
8. CIOCCO A: Variation in the sex ratio at birth in the U.S. *Hum Biol* 10:36–64, 1938
9. LEJEUNE J, TURPIN R: L'influence de l'âge des parents sur la masculinité de naissances vivantes. *CR Acad Sci [D]* (Paris) 244:1833–1835, 1957
10. NOVITSKI E: The dependence of the secondary sex ratio in humans on the age of the father. *Science* 117:531–533, 1953
11. BERNSTEIN ME: Parental age and the sex ratio. *Science* 118:448–449, 1953
12. NOVITSKI E, SANDLER L: The relationship between parental age, birth order and the secondary sex ratio in humans. *Ann Hum Genet* 21:318–324, 1956
13. MORAN PAP, NOVITSKI E, NOVITSKI C: Paternal age and the secondary sex ratio in humans. *Ann Hum Genet* 32:315–317, 1969
14. POLLARD GN: Factors influencing the sex ratio at birth in Australia, 1902–65. *J Biosocial Sci* 1:125–144, 1969
15. NOVITSKI E, KIMBALL AW: Birth order, parental ages, and sex of offspring. *Amer J Hum Genet* 10:268–275, 1958
16. MANTEL N, HAENSZEL W: Statistical aspects of the analysis of data from retrospective studies of disease. *J Nat Cancer Inst* 22:720–748, 1959
17. MANTEL N: Chi-square tests with one degree of freedom: extensions of the Mantel-Haenszel procedure. *J Amer Statistical Assn* 58:690–700, 1963
18. STARK CR, MANTEL N: Effects of maternal age and birth order on the risk of mongolism and leukemia. *J Nat Cancer Inst* 37:687–698, 1966
19. VISARIA PM: Sex ratio at birth in territories with a relatively complete registration. *Eugen Quart* 14:132–142, 1967
20. GOODMAN LA: Some possible effects of birth control on the human sex ratio. *Ann Hum Genet* 25:75–81, 1961; also in *Readings in Mathematical Social Science*, edited by LAZARSFELD PF, HENRY NW, Chicago, Science Research Associates, 1966, pp 311–317
21. KIRBY DRS, McWHIRTER KG, TEITELBAUM MS, et al: A possible immunological influence on sex ratio. *Lancet* 2:139–140, 1967